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# **Food Prices in Six Developing Countries and the Grains Price Spike**

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## **Abstract**

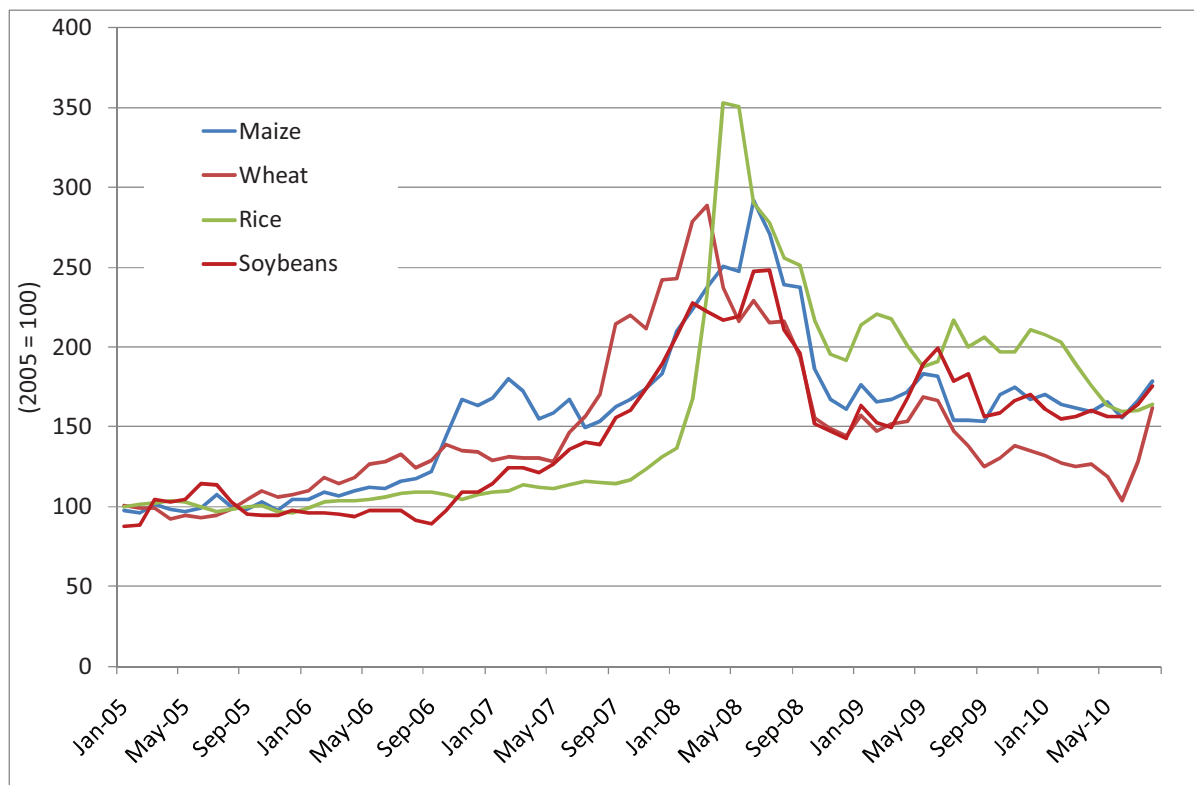
The paper analyzes the pass-through of world maize and rice prices in six developing countries (Benin, Kenya, Malawi, Nepal, Peru and Vietnam) both at the national and sub-national (regional) levels with the view of considering the relevance of world prices for national prices (high for maize, low for rice) and the representativeness of national average prices for prices throughout the country (high in Kenya and Malawi, low in Peru). The variability of prices across regions generally, but not invariably, increases when prices are high. Kenya and Vietnam have been relatively successful and Malawi least successful in insulating consumers from volatility in world prices.

Department of Economics, University of Trento, via Inama 5, 38122 Trento, Italy; email: [cgilbert@economia.unitn.it](mailto:cgilbert@economia.unitn.it) .

This paper has been prepared for the workshop “Impacts, Responses and Initial Lessons of the Financial Crisis for Low Income Countries”, organized by the Danish Institute for International Studies, Copenhagen 14-15 October 2010. It is based on a study examining the impact on the agricultural terms of trade of the energy and grain price surge on a number of developing and less developed countries undertaken by the author for the Food and Agriculture Organization of the United Nations (FAO). I am grateful to the following collaborators in the FAO project: Rose Edwige Fiamoh (Benin), Harriet Mugeru (Kenya), Sridhar Thapa (Nepal), Hien Minh Vu (Vietnam), Santos Maza Ysillupu (Peru) and Wouter Zant (Malawi). A preliminary version of the paper was presented at the BREAD Summer Workshop in Development Economics, Alba di Canazei, 30 July 2010. I am grateful to Alain de Janvry, Elisabeth Sadoulet and other participants at the workshop for useful comments. The views expressed are those of the author and not the FAO.

## 1 Introduction

The prices of food commodities on world markets increased dramatically in 2007-08. Subsequently, in 2009, prices fell back although, except for wheat, not to their pre-spike levels – see Figure 1 which charts the world prices of the four principal grains, maize (corn), wheat, rice and soybeans.<sup>1</sup> The summer of 2010 has witnessed catastrophic weather conditions in much of the northern hemisphere resulting in renewed upward pressure on the wheat price which nevertheless remains well below its 2008 peak. These developments have resulted in a widespread view that the combination of rapid economic growth in much of Asia with more variable weather conditions perhaps in part caused by global warming will result in higher and more variable food prices over at least the next decade.



**Figure 1-1: International grains prices, 2005-10**

I discussed the causes of the 2007-08 price spike in Gilbert (2010a) where I argued that demand-side factors were predominant. I also suggested that the evidence was weak that biofuels feedstock demand for food commodities played a major role as had been suggested by Mitchell (2008). Instead, I indicated that futures market factors may have amplified movements generated by rapid economic growth and loose monetary policies – see also Mayer (2009) and Gilbert (2010b). Consistent with this view, prices collapsed rapidly from September 2008 with the onset of the financial crisis. If these arguments are correct, the

<sup>1</sup> Source: IMF, *International Financial Statistics*.

2007-08 may be seen as one of the symptoms of the general financial latitude that resulted in the 2008 financial crisis, and the collapse in food prices as a consequence of the financial crisis itself.

Futures trading remains relatively unimportant in the rice market. Rice futures are traded in Bangkok and Chicago but volumes are small and speculative interest slight. Christiaensen (2009) has emphasized the role of export restrictions, imposed by major exporting countries, in extending the wheat and maize price spike to rice – see also World Bank (2007).

The summer 2010 price movements are both similar and different to the 2007-08 spike. Differently from 2007-08, these movements originated in weather-induced harvest losses, in particular in Russia. Again differently from 2008, the sharp jump in price has been confined to wheat with only modest rises in other grains prices.<sup>2</sup> Similarly to 2007-08, these price movements have been amplified by futures markets and by the imposition of an export ban by an important exporting country, in this case Russia.

It remains unclear whether or not this implies that food prices will be inexorably either higher or more volatile in the future. Over the long period, agricultural prices have fallen in real terms as the result of steady improvements in agricultural productivity. There is every reason to expect this process to continue and there is little strong evidence that the Ricardian margin is likely to reduce yields at the margin. Turning to volatility, Gilbert and Morgan (2010) argue that this is determined by the interaction of production (harvest) and consumption shocks with supply and demand elasticities. Increased weather variability may therefore translate into higher food price volatility but volatility remains lower over recent years than in the 1970s, the previous decade associated with high food prices.

Food forms a relatively small proportion of household budgets in the developed world. High or volatile food prices remain a concern for poor households, and hence also for governments, but overall, high energy prices have probably been more problematic than high food prices. In the developing world, by contrast, food can be the principal expenditure item for a large proportion of the population. High and volatile food prices therefore have a much more substantial impact on welfare levels. Furthermore, changes in the prices of food commodities impact the rural-urban terms of trade and, for countries which are dependent on food imports, on the trade balance, reserves and the exchange rate. This raises the important issue of the extent to which high and volatile prices on world markets are passed through into developing countries.

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<sup>2</sup> The August-June price rise in wheat was 56.1%. The comparable rises for the other three grains considered in Figure 1 were 14.9% for maize, 12.0% for soybeans and 2.9% for rice.

There are three groups of factors which may limit pass-through.

- a) Governments may run successful stabilization policies serving to insulate domestic consumers from changes in world prices. This has been true of many important Asian rice producing-consuming countries (Dawe, 2007).
- b) High transport costs may insulate countries, or regions within countries, from part of all of movements in world prices. These costs define a band between the low export parity price, at which it is profitable for the country to export, and the high import parity price, at which imports become profitable. The band moves up and down with world prices but within the band, the commodity will neither be imported nor exported and will vary largely independently of world prices. This can result in very high levels of volatility of food prices in landlocked countries – see Dana *et al* (2006) on Malawi and Zambia and Daviron (2008) on francophone Africa.
- c) The prices which we regard as measuring world prices may be unrepresentative of actual transactions prices in world trade. I shall suggest below that this is the case in the rice market. For other food commodities, such as cassava, there is no recognized international price.

The objective of this paper is to throw further light on the evolution of important food prices in a group of developing countries – Benin, Kenya, Malawi Nepal, Peru and Vietnam. In this paper, which draws on a larger study, I focus on maize and rice. White maize is the staple food crop in most of southern and eastern Africa. Yellow maize is an important animal feedstock in Europe and North America but is also directly consumed by humans although not in huge quantities. However, it is a major food in parts of central America. It has become a major biofuel feedstock in the United States. Rice is the major food staple throughout most of Asia where maize is unimportant in human consumption. It is also a significant source of carbohydrate throughout the developing world even where it is not the staple food.

Here, I look at two issues. The first is that of pass-through and the relevance, if any, of world prices for the evolution of local prices (section 2). The second is that of regional price variability within the countries (section 3). The same issue has been analyzed for maize, rice, soybeans and wheat in Guatemala by de Janvry and Sadoulet (2010). Daviron (2008) looked at transmission of the high 2007-08 wheat and rice prices in six African countries.<sup>3</sup> In both cases, the analysis finishes in mid 2008 so the authors unable to examine the effects of the September 2008 financial crisis on falling prices.

## **2 World prices and pass-through**

The prices considered as world prices are prices set on major international futures markets. They may either be spot prices or, more generally, the price of the delivery contract. For

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<sup>3</sup> Cameroon, Guinea, Madagascar, Mali, Niger and Senegal.



maize, the standard reference price is the Chicago Board of Trade (CBOT)<sup>4</sup> corn (i.e. yellow maize) futures price. However, for the African countries (Kenya and Malawi), it seems possible that the South African Futures Exchange (SAFEX) white maize futures price may be relevant. For rice, the standard reference price is the Bangkok spot price reflecting the fact that Thailand is the major world rice exporter. This spot price is related to the white rice futures contract on the Agricultural Futures Exchange of Thailand (AFET) although trading in this contract remains very thin. Rough (i.e. unmilled) rice is also traded in Chicago on CBOT but volumes are low relative to those on other Chicago grains markets and the contract is regarded as being primarily of domestic interest. The first question is therefore the relevance of these international prices to prices in developing countries.

Supposing one or more world prices is relevant in a particular country, the second question is the rapidity of adjustment of local prices to the world price. This literature originated with Timmer *et al.* (1983) and Mundlack and Larson (1992) who regressed local on world prices. This procedure is problematic when prices trend, or are more generally non-stationary. Estimation in first differences, as in de Janvry and Sadoulet (2010), results in estimates which are robust with respect to non-stationarity but which may only measure short term responses. This may result in under-estimation of adjustment if adjustment is slow. Baffes and Gardner (2003) used the error correction specification to estimate pass-through. This gives rise to estimates of both impact and equilibrium responses. Error correction can be justified by the Granger Representation Theorem (Engle and Granger, 1987) if the two prices considered are cointegrated but that hypothesis is not directly tested. When three or more prices are considered, as in this paper, the supposition that there is a single cointegrating vector is problematic.

Following Ardeni (1989), it has become standard to adopt cointegration-based methods following Johansen (1988). This is the approach adopted here. It has four advantages over more traditional techniques

- a) The number of cointegrating vectors is determined by the data.
- b) Short run adjustment responses are distinguished from equilibrium outcomes (if present).
- c) The equilibrium outcome is not restricted to be unity.
- d) Adjustment of national and world prices is considered symmetrically allowing the possibility of reverse pass-through from the former to the latter.

Baulch (1987) and Barrett and Li (2002) criticized the cointegration approach as failing to allow for the wedge between export and import parity prices. They prefer to adopt a switching regime model. By contrast, Rapsomanikis and Karfakis (2008) use the Balke and Formby (1997) threshold cointegration model to accomplish the same objective. The analysis in this paper relies on relatively short time series which makes the more

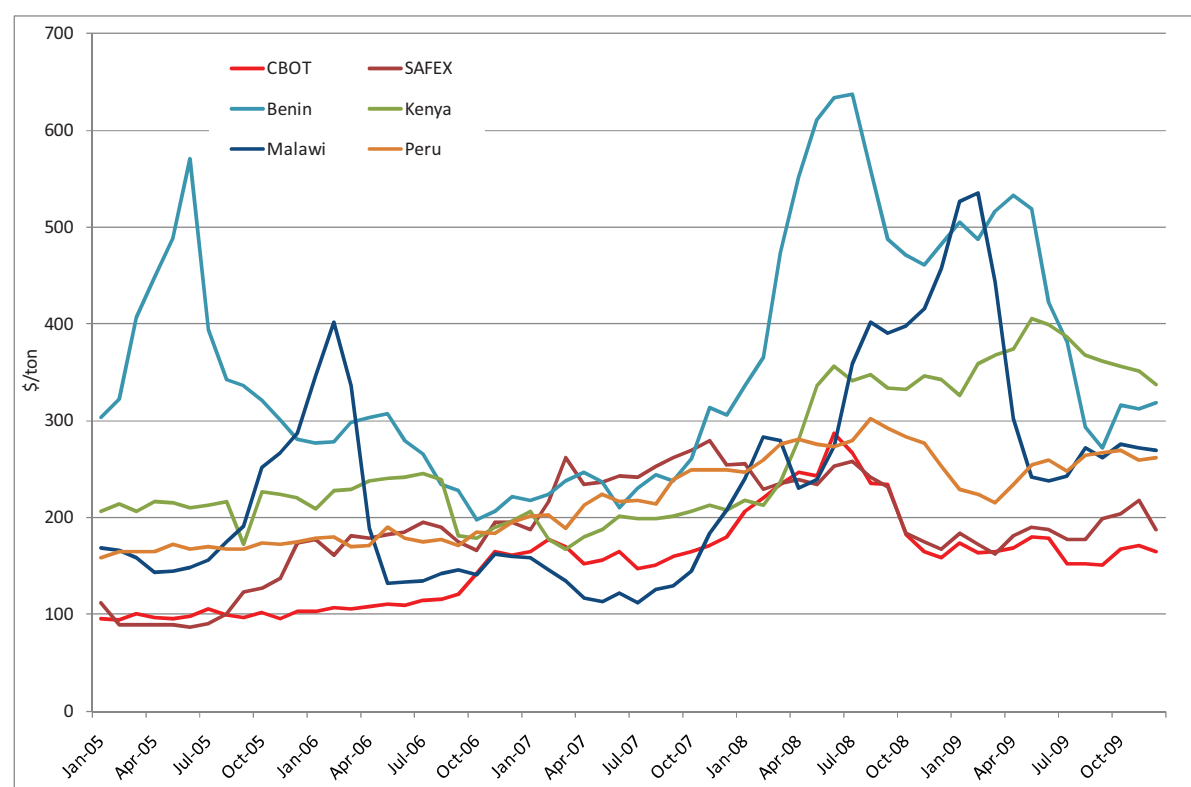
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<sup>4</sup> CBOT is now part of the Chicago Mercantile Exchange (CME) group.

sophisticated post-Ardeni methods unattractive. Balke and Formby (1997) report that standard cointegration test procedures, such as the Johansen (1988) VAR-based test, work reasonably well if the true process has the threshold structure with reversion taking place only beyond the (in this case export and import parity) thresholds. On this basis, I am happy to reply on the Johansen procedure.

## 2.1 Maize

Figure 2 charts maize prices from January 2005 to December 2010 at the national level for the three African countries under consideration and Peru. These national prices are medians of prices across a range of locations – see sections 3.1 and 3.2 below. In addition, the figure shows the South African (SAFEX) and U.S. (Chicago) free market prices. All prices have been converted to US dollars per ton at the prevailing exchange rate. The two international prices move closely together but the other four prices only track these approximately.



**Figure 2: US dollar maize prices, 2005-09**

Table 1 lists the nominal and real price ranges in each of the four countries considered (columns 3 and 4) and also the 2005-09 price change (columns 1 and 2). The range measures the maximum extent of the price spike while the change shows the long run impact, if any. The final two columns give the standard deviations of price changes over the period. Care must be taken in the interpretation of these real prices since, in countries in which maize forms a substantial component of the household budget, deflation makes little sense for poor households.

<b>Table 1</b> <b>Maize price changes, 2005-2009</b>						
	January 2005 - December 2009		Price range over the same period		Standard deviation of monthly changes	
	Nominal	Real	Nominal	Real	Nominal	Real
Free market – Chicago	71.5%	50.8%	204.3%	140.7%	26.1%	24.3%
Free market - SAFEX	67.2%	47.1%	223.9%	178.3%	31.3%	29.3%
Benin	5.0%	-18.4%	223.3%	217.5%	38.7%	37.0%
Kenya	62.9%	-14.2%	141.8%	96.3%	27.7%	27.9%
Malawi	59.2%	36.3%	380.6%	288.2%	52.3%	49.4%
Peru	65.3%	27.7%	91.1%	51.5%	16.5%	15.1%
Average price	41.7%	24.6%	130.4%	126.6%	20.0%	20.3%
The first two columns of the table gives the percentage change in the free market (Chicago and SAFEX) maize prices and local rice prices respectively converted to US dollars and local prices deflated by the local CPI (national prices) or the Advanced Countries Export Unit Values (exchange prices and average price) over the period January 2005 – December 2009. The second two columns give the percentage range between the maximum and minimum prices over the same period and the final two columns the standard deviation of logarithmic price changes on an annual basis. <sup>5</sup> Source for exchange rates, CPI and export unit value indices: IMF, <i>International Financial Statistics</i> .						

The maize price range for Malawi exceeds that in world markets while those for Kenya and Peru are much more modest. Over the entire period, price rises are comparable to those in world markets except in Benin. Maize prices in Benin were very high in 2005 for local reasons and hence the 2005-09 price rise appears misleadingly modest – see Figure 2. A 2007 base would have given very different results. Effective price stabilization has resulted in relatively stable prices in Kenya and, to the extent that this interpretation can be sustained, the maize price has declined in real terms. It is evident that that maize price variability has been acute in Malawi, whatever basis is used for making the judgement.

Statistical analysis of the nominal dollar maize price series over the five years 2005-09 confirms that the series are non-stationary, although this result is quite marginal for the SAFEX series (DF = -2.74 against a 5% critical value of -2.91). This near stationarity precludes the use of cointegration analysis to analyze the inter-relationship with the series using the Johansen procedure on a pair-wise basis for each of the Beninois, Kenyan, Malawian and Peruvian (logarithmic) prices and each of the log exchange prices gives within a VAR(2), fails to reject  $\text{rank}(\alpha\beta') = 0$  implying each of the series is stationary. This suggests moving to longer series, I therefore analyze data from 1999-2009 for the three exchange prices and for Kenya, Malawi and Peru (price starts in 2000), for which long price series are available.<sup>6</sup>

<sup>5</sup> i.e. the standard deviation of monthly price changes annualized by multiplying by  $\sqrt{12}$

<sup>6</sup> These prices are the national prices as stated by government departments or international agencies: Kenya – Ministry of Agriculture; Malawi – World Bank; Peru – Ministry of Agriculture. Prices converted to U.S. dollars per kilogram at prevailing exchange rates; source: IMF, *International Financial Statistics*.

Figure 3 charts these five series. The Malawian price series shows three “hungry season” peaks corresponding to the poor harvests in 2001 and 2005 and to the spike in world prices in 2008. The Kenyan series shows much less variation prior to the 2007-08 spike but prices appear to have been generally higher than those in Malawi.

ADF tests confirm that the four (logarithmic) series are all non-stationary, although this result is marginal for the Malawian series – see the initial column of Table 2. This allows application of the Johansen procedure to analyze cointegration. I first consider the two exchange prices. A preliminary check establishes that I can consider this within a VAR(2) framework. I fail to reject the hypothesis that  $rank(\alpha\beta') \leq 1$  but reject the hypothesis that  $rank(\alpha\beta') = 0$  confirming that there is a single cointegrating vector – see the second and

third column of Table 2. The estimated  $\alpha$  coefficients are  $\begin{pmatrix} \hat{\alpha}_{Chicago} \\ \hat{\alpha}_{SAFEX} \end{pmatrix} = \begin{pmatrix} -0.047 \\ 0.087 \end{pmatrix}$  with standard

errors  $\begin{pmatrix} 0.023 \\ 0.034 \end{pmatrix}$ . Both coefficients differ significantly from zero implying each market reacts

to the other, but the coefficient for SAFEX is approximately double that for the Chicago market consistent with the leading role played by Chicago. I also fail to reject the hypothesis that this is a unit cointegrating vector ( $\chi^2_1 = 0.46$  with p-value 50.0%) implying that over the long term the Chicago and SAFEX prices may be seen as moving in line with each other.

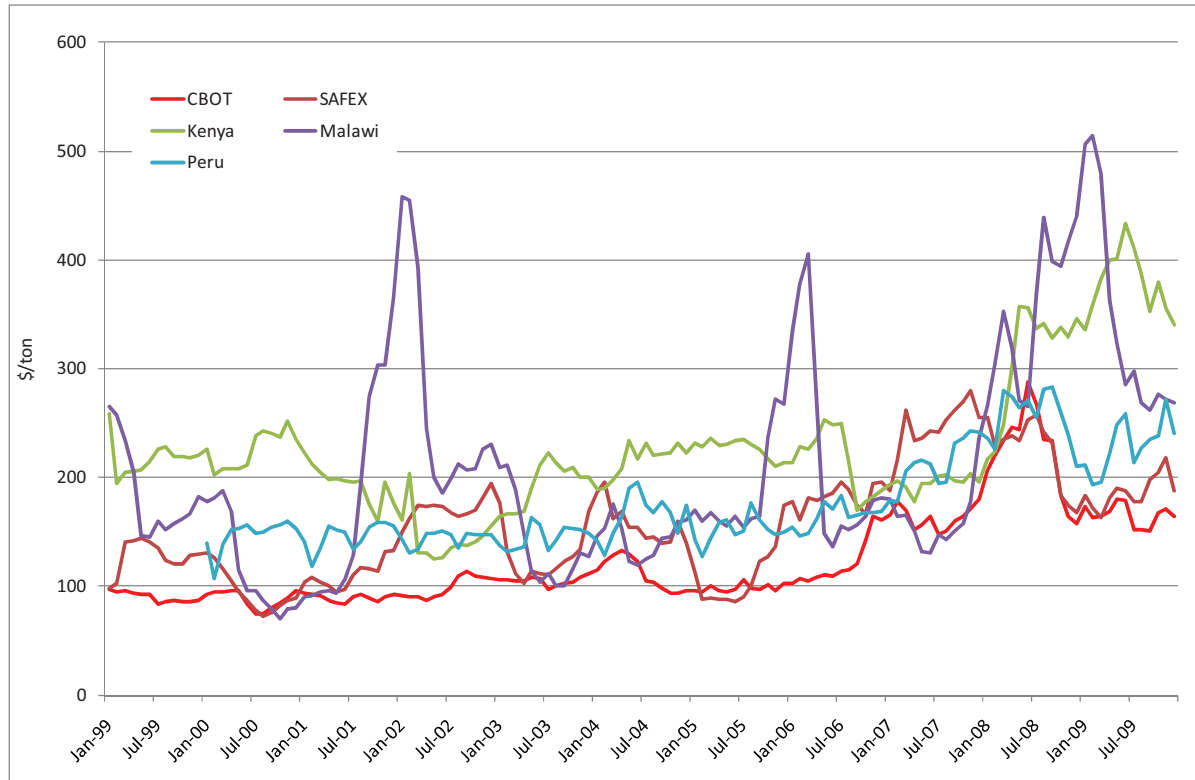


Figure 3: US dollar maize prices, 1999-2009

I now add the Kenyan, Malawian and Peruvian prices in turn to the VAR. Given that we already know that the Chicago and SAFEX prices are cointegrated, cointegration of the African prices requires  $rank(\alpha\beta')=2$ . This result is established for Malawi and Peru but not for Kenya – see Table 2. Despite its proneness to weather-related shortages, in the long term, the Malawian maize market appears integrated with world markets. In line with the visual impression obtained from Figure 3, this is not true of Kenya.

It is interesting to look at the Malawian and Peruvian cases in greater detail. Since  $rank(\alpha\beta')=2$ , we have only identified a two dimensional basis for the space in which the cointegrating vectors lie. It is therefore open to us to rotate the estimated cointegrating vectors within that space. Since we have already established that the Chicago and SAFEX

prices are cointegrated with unit cointegrating vector, we can reasonably impose  $\beta_1 = \begin{pmatrix} 1 \\ -1 \\ 0 \end{pmatrix}$

with  $\alpha_{13} = 0$  (i.e. the two local prices do not react to temporary discrepancies between SAFEX and Chicago). Thereafter, the procedure differs between the two countries.

<b>Table 2</b> <b>Statistical Properties of Maize Price Series, 1999 - 2009</b>					
	Stationarity	Trace cointegration tests			Implied # cointegrating vectors
		rank 0	rank $\leq 1$	rank $\leq 2$	
Chicago	ADF(1) - 1.36	18.7	2.23	-	1
SAFEX	ADF(1) - 2.16	[1.4%]	[13.6%]		
Kenya	DF - 1.22	37.8	10.7	2.51	1
		[0.4%]	[23.6%]	[11.3%]	
Malawi	ADF(2) - 2.58	34.9	16.7	2.46	2
		[1.1%]	[3.1%]	[11.7%]	
Peru	ADF(2) - 2.37	44.3	15.5	1.94	2
		[< 0.1%]	[4.8%]	[16.4%]	
Sample April 1999 (Peru: April 2000) – December 2009. The ADF lag was selected over the range 0-3 on the basis of the Akaike Information Criterion. 5% critical value is – 2.91. Cointegration is examined using the Johansen trace tests within a set of bivariate VAR(2) models. The combined first two rows consider bivariate cointegration between the Chicago and SAFEX price. The third and fourth row consider trivariate cointegration between Chicago, SAFEX and the Kenyan and Malawian prices in turn. The reported statistics test $rank(\alpha\beta') \leq r$ ( $r=0,1,2$ ). Tail probabilities in “[.]” parentheses.					

Malawi: Normalizing the second cointegrating vector, I hypothesize that this depends

equally on the two exchange prices so  $\beta_2 = \begin{pmatrix} \frac{1}{2} \\ \frac{1}{2} \\ -1 \end{pmatrix}$ . Since it is reasonable to suppose that the

Chicago price is unaffected by the maize situation in Malawi, we can impose  $\alpha_{21} = 0$  in the Malawian case. More tendentiously, we can impose the same condition on the SAFEX price, i.e.  $\alpha_{22} = 0$ . This implies a total of 5 restrictions on the  $\alpha\beta'$  matrix. The likelihood ratio fails to reject these restrictions ( $\chi^2_5 = 2.39$  with p-value 79.2%) implying an acceptable identification.

The estimated  $\alpha$  matrix is  $\begin{pmatrix} \hat{\alpha}_{Chicago} \\ \hat{\alpha}_{SAFEX} \\ \hat{\alpha}_{Malawi} \end{pmatrix}' = \begin{pmatrix} -0.056 & 0.000 \\ 0.087 & 0.000 \\ 0.000 & 0.104 \end{pmatrix}$  with standard errors  $\begin{pmatrix} 0.026 & - \\ 0.038 & - \\ - & 0.028 \end{pmatrix}$ .

Conditional on the acceptability of these restrictions, the estimates show that we can reject both the null hypotheses  $\alpha_{Chicago} = 0$  and  $\alpha_{SAFEX} = 0$  implying that the markets are inter-dependent. However, as one might expect, SAFEX reacts more than Chicago to deviations from parity.

Peru: It seems unlikely that the South African white maize price can influence the yellow

maize price in Peru so I set  $\beta_2 = \begin{pmatrix} 1 \\ 0 \\ -1 \end{pmatrix}$ . As in the case of Malawi, we can use a “small

country” restriction to suppose that Peruvian prices have no influence on the exchange prices allowing us to set  $\alpha_{21} = \alpha_{22} = 0$ . As again in Malawi, the likelihood ratio fails to reject these restrictions ( $\chi^2_4 = 1.02$  with p-value 90.7%) implying an acceptable identification. The

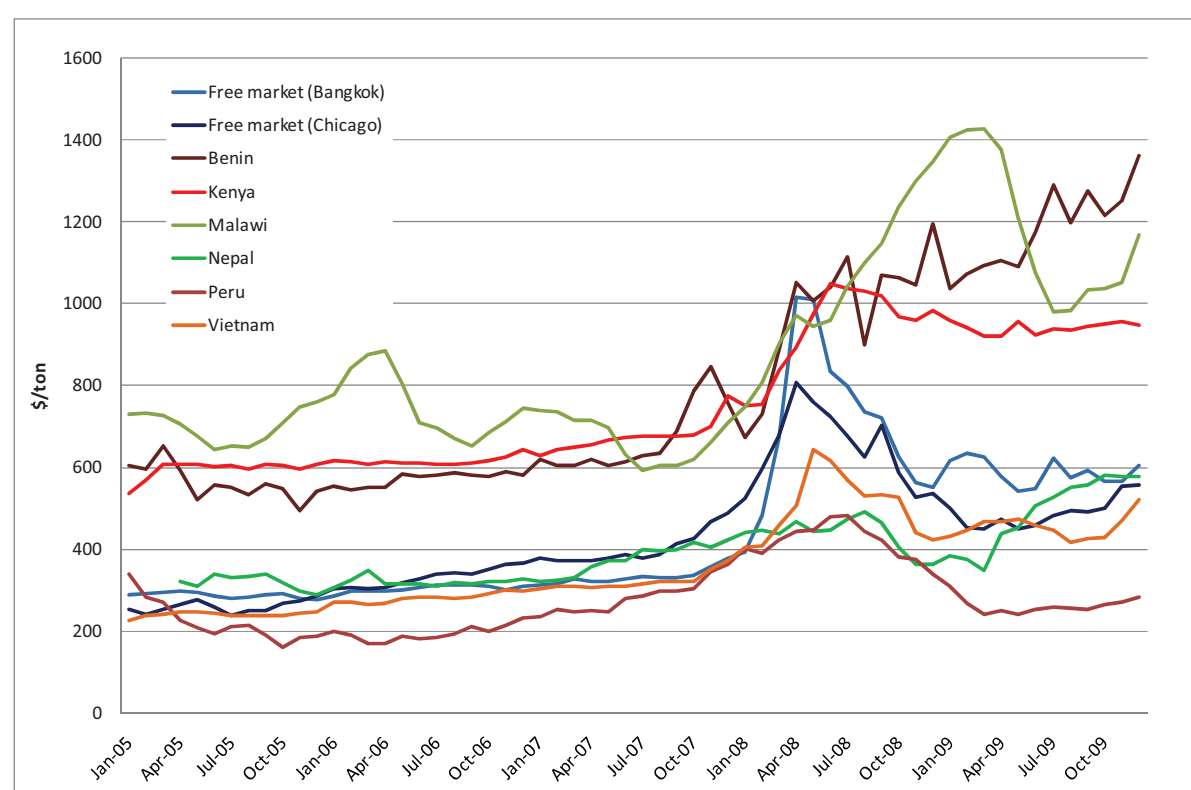
estimated  $\alpha$  matrix is now  $\begin{pmatrix} \hat{\alpha}_{Chicago} \\ \hat{\alpha}_{SAFEX} \\ \hat{\alpha}_{Malawi} \end{pmatrix}' = \begin{pmatrix} -0.062 & 0.000 \\ 0.074 & 0.000 \\ 0.000 & 0.409 \end{pmatrix}$  with standard errors  $\begin{pmatrix} 0.026 & - \\ 0.039 & - \\ - & 0.070 \end{pmatrix}$ . It

is reassuring to see that the estimated  $\alpha$  coefficients for the two exchange prices are very similar in the Malawian and Peruvian models.

Both the Kenyan and Malawian governments actively intervene in their respective maize markets. In Kenya, the National Cereals and Produce Board (NCPB) purchases between 10% and 20% of the domestically marketed maize output in Kenya, mainly from large scale farmers. In Malawi, the Agricultural Development and Marketing Corporation (ADMARC) sets prices for maize purchase which have often been below the export parity price. The results reported above and in Table 1 indicate that while Kenya has been relatively successful in stabilizing its domestic maize prices, Malawi has been much less so. This may reflect the much more frequent Malawian dependence on maize imports.

## 2.2 Rice

Figure 4 charts local rice prices from January 2005 to December 2010 for the six countries under consideration, and in addition includes the world free market (Bangkok and Chicago) prices.<sup>7</sup> All prices have been converted to US dollars per ton at the prevailing exchange rate. The figure shows the prices falling into two groups: three low price countries (Nepal, Peru and Vietnam) and the three high price African countries (Benin, Kenya and Malawi). The world free market price is typically closer to prices in the non-African group but rose to African levels in 2008. It is apparent from Figure 4 that the Bangkok price has been much more variable than any of the national prices. This is shown in Table 3. Although the change in the Bangkok price over the five year period is comparable to that in the six countries considered, the range between the maximum and minimum prices was approximately double for the former.



**Figure 4: US dollar rice prices, 2005-09**

Table 3 shows considerable variation across countries in the evolution of deflated rice prices. As was the case of maize, care must be taken in the interpretation of these real prices in countries in which rice forms a substantial component of the household budget.

<sup>7</sup> The Chicago price relates to rough rice whereas the remaining prices are for white rice. I have used a conversion factor of 1.67 to convert the Chicago price onto a milled basis. Rice prices for Malawi, Nepal, Peru and Vietnam are calculated as the medians of prices across localities - see section 3.3. Local prices are not available for Benin and Kenya: Benin – “regular rice”, Cotonou, urban market, F.CFA/kg, Ministère du commerce, Direction de la Promotion du Commerce Intérieur; Kenya – Average wholesale price, rice grade 2, loose, Ksh/kg, Kenyan National Bureau of Statistics. Because of different specifications, prices are not perfectly comparable across countries.



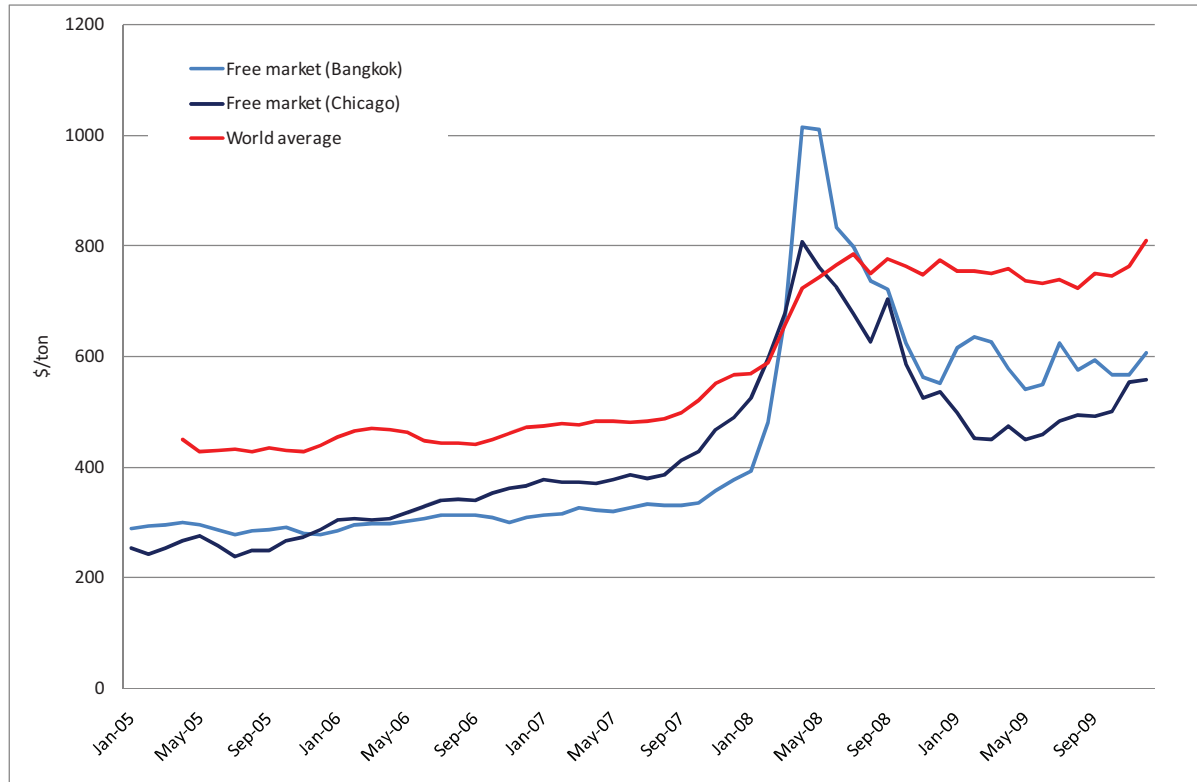
With this qualification, the statistics show a decline in real rice prices in Peru over the five years 2005-09 and almost no change in Kenya, while real prices have risen by 20%-40% in Malawi and Nepal, and 60%-75% on Benin and Vietnam.

<b>Table 3</b>						
<b>Rice price changes, 2005-2009</b>						
	January 2005 - December 2009		Price range over the same period		Standard deviation of monthly changes	
	Nominal	Real	Nominal	Real	Nominal	Real
Free market (Bangkok)	93.3%	75.9%	223.9%	174.3%	39.4%	37.0%
Free market (Chicago)	47.5%	34.3%	118.0%	87.9%	25.7%	22.2%
Benin	124.9%	74.6%	175.7%	99.9%	27.0%	25.1%
Kenya	76.5%	-7.1%	95.2%	33.3%	10.7%	9.6%
Malawi	59.9%	36.8%	140.3%	91.6%	19.3%	16.9%
Nepal	79.3%	23.5%	100.6%	42.7%	19.8%	15.8%
Peru	-16.7%	-35.6%	202.4%	135.1%	27.4%	26.2%
Vietnam	129.7%	58.1%	184.1%	92.6%	19.4%	18.2%
Average price	80.1%	85.3%	89.2%	102.7%	10.1%	11.6%
The first two columns of the table gives the percentage change in the free market rice prices and local rice prices respectively converted to US dollars and local prices deflated by the local CPI over the period January 2005 (Nepal and Average: April 2005) – December 2009. The second two columns give the percentage range between the maximum and minimum prices over the same period and the final two columns the standard deviation of logarithmic price changes (at an annual rate). The deflator for the free market price and the average price is Advanced Countries Export Unit Values. Source for exchange rates, CPI and export unit value indices: IMF, <i>International Financial Statistics</i> .						

Prices in each country clearly reflect specific local conditions. This motivates consideration of the average world price, averaging over the six countries considered. Figure 5, which compares this average price with the free market price, confirms this greater variability of the free market prices but also shows the general tendency for rice prices to rise over the entire five year period.

Table 4 lists the statistical properties of the nine nominal U.S. dollar rice price series considered in Table 3. ADF tests show that all nine logarithmic series are non-stationary. Table 4 also reports Johansen cointegration tests. In the first two rows, I ask whether the Bangkok and Chicago free market prices are cointegrated. The results are inconclusive but consistent with a lack of cointegration – the Johansen test marginally fails to reject  $\text{rank}(\alpha\beta') \leq 1$  which would imply both prices are stationary, contradicting the clear results from the ADF tests. Ignoring that result, the second test also fails to reject  $\text{rank}(\alpha\beta') = 0$ . I conclude that the rank is zero and there is no cointegration. Figure 5 shows the Chicago price trending upward relative to Bangkok over the five year period. (The ADF test on the logarithmic difference of the two prices gives a statistic of  $\text{ADF}(1) = -2.26$ ).





**Figure 5: Free market and average world rice price, 2005-09**

To check on this result, the test was rerun using the longer sample March 1999 – December 2010. ADF tests again confirm non-stationarity (Bangkok,  $ADF(2) = -2.58$ ; Chicago,  $ADF(1) = -0.76$ ). The Johansen test now gives a clear rank zero result. The tests fail to reject both  $\text{rank}(\alpha\beta') \leq 1$  (trace statistic 0.50 with p-value 48.1%) and  $\text{rank}(\alpha\beta') = 0$  (trace statistic 9.74 with p-value 30.7%). I conclude that the two rice exchange prices are indeed not cointegrated.

I now return to the original 2005-09 sample and add the six national rice prices, one at a time, to give a trivariate VAR(2). Whereas in the case of maize, where the two exchange prices are cointegrated, I tested for  $\text{rank}(\alpha\beta') = 2$ , in rice, lacking this cointegration, I test for  $\text{rank}(\alpha\beta') = 1$ .

A cointegrating rank of one is established for Peru. For Vietnam, the failure to reject  $\text{rank}(\alpha\beta') = 0$  is marginal. The test outcomes are problematic for Kenya and Malawi where the test rejects  $\text{rank}(\alpha\beta') \leq 1$  (marginally in the Kenyan case) which would imply all three prices are stationary, again contradicting the results of the ADF tests. However, if we override this result, we again establish a cointegrating rank of unity. By contrast, cointegration is completely rejected in the case of Benin and Nepal where local prices continued to rise through 2009 while they fell back elsewhere – see Figure 4. The price

constructed as the average of the six national prices does appear to be cointegrated with the two free market prices.<sup>8</sup>

Table 4 Statistical Properties of Rice Price Series, 2005-09					
	Stationarity	Cointegration tests			Implied # cointegrating vectors
		rank 0	rank $\leq 1$	rank $\leq 2$	
Free market (Bangkok)	ADF(2) -1.12	12.8	3.94	-	0 (?)
Free market (Chicago)	ADF(1) -1.30	[12.3%]	[4.7%]		
Benin	ADF(2) 0.43	24.2 [19.9%]	6.59 [63.1%]	0.31 [57.5%]	0
Kenya	ADF(1) - 0.64	43.1 [0.1%]	11.2 [21.5%]	4.00 [4.6%]	1 (?)
Malawi	ADF(3) - 1.36	32.5 [2.3%]	12.7 [12.8%]	5.56 [1.8%]	1 (?)
Nepal	DF - 0.23	25.2 [15.8%]	5.80 [72.1%]	0.27 [60.2%]	0
Peru	ADF(1) - 1.19	31.8 [2.8%]	11.2 [20.3%]	3.70 [5.4%]	1
Vietnam	ADF(1) - 0.89	28.6 [7.0%]	7.68 [30.6%]	2.28 [13.1%]	0 or 1
Average price	ADF(1) - 0.22	31.9 [2.8%]	9.37 [33.9%]	1.52 [21.8%]	1
Sample: March 2005 (rank tests; Nepal and Average – June 2005) and May 2005 (ADF tests; Nepal and Average – August 2005) – December 2009. The ADF lag was selected over the range 0-3 on the basis of the Akaike Information Criterion. 5% critical value is – 2.91. Cointegration is examined using the Johansen trace tests within a set of bivariate VAR(2) models considering the free market price with each of the other prices. The reported statistic tests $\text{rank}(\alpha\beta') \leq 1$ . Tail probabilities in “[.]” parentheses.					

The results reported in Table 4 demonstrate possible links between the various national prices, with the exception of those of Benin and Nepal,<sup>9</sup> and the two world prices. To enable examination of these links in greater detail, in Table 5 I report the estimated  $\alpha$  and  $\beta$  coefficients.<sup>10</sup> The first two columns of the table test whether the  $\beta$  coefficients for respectively the Bangkok and Chicago prices can be set to zero. I fail to reject the hypothesis with respect to the Chicago price for Kenya and, marginally, for Vietnam (column 2). The same hypothesis with respect to the Bangkok price is rejected for all four countries (column 1). The third and fourth columns of the table report the estimated  $\beta$  coefficients, relative to

<sup>8</sup> This result is also problematic. If any one of the six prices composing the average is not cointegrated with the exchange prices, the average itself cannot be cointegrated. However, using relative short samples, apparently contradictory results of this sort can emerge.

<sup>9</sup> Agricultural prices in Nepal relate to Indian prices in the first instance, and not to world prices.

<sup>10</sup> Benin and Nepal are excluded because of lack of cointegration.

a normalization of minus one on the coefficient of the national price. I set the coefficient  $\beta_{Ch}$  on the Chicago price to zero for Kenya and Vietnam based on the test outcomes reported in column 2 and also for Malawi where  $\beta_{Ch}$  is estimated as negative. Similarly, I set  $\beta_{Bk} = 0$  for Peru where this coefficient was estimated as negative. In summary, the Kenyan, Malawian and Vietnamese rice prices are seen as being cointegrated with the Bangkok price and the Peruvian price with the Chicago price. The final three columns give the estimated  $\alpha$  coefficients. The size of these coefficients  $\alpha_j$  shows the speed at which the national prices react to any discrepancy with respect to world prices. Reaction is seen as fastest in Vietnam and slowest in Malawi.

<b>Table 5</b> <b>Estimated <math>\alpha</math> and <math>\beta</math> vectors</b>							
	Tests of $\beta$ vector ( $\chi^2_{11}$ )		Estimated $\beta$ vector		Estimated $\alpha$ vector		
	$H_0: \beta_{Bk} = 0$	$H_0: \beta_{Ch} = 0$	$\beta_{Bk}$	$\beta_{Ch}$	$\alpha_{Bk}$	$\alpha_{Ch}$	$\alpha_j$
Kenya	22.6 [< 0.1%]	0.02 [89.0%]	0.585 (0.024)	-	-0.611 (0.167)	-0.342 (0.139)	0.181 (0.060)
Malawi	9.96 [1.6%]	4.77 [2.9%]	0.611 (0.107)	-	-0.076 (0.066)	0.003 (0.053)	0.106 (0.034)
Peru	9.19 [2.4%]	13.1 [< 0.1%]	-	1.038 (0.166)	-0.074 (0.057)	-0.077 (0.045)	0.125 (0.054)
Vietnam	15.1 [0.1%]	3.55 [6.0%]	0.821 (0.051)	-	-0.422 (0.113)	-0.245 (0.096)	0.059 (0.063)
Average	12.3 [0.5%]	0.47 [49.1%]	0.738 (0.049)	-	-0.371 (0.114)	-0.207 (0.093)	0.024 (0.041)
The table reports the estimated $\alpha$ and $\beta$ vectors for the reduce rank Johansen estimates of rank one trivariate VAR(2) models. Suffices <i>Bk</i> and <i>Ch</i> refer to the Bangkok and Chicago prices; the suffix <i>j</i> refers to the row country's national price. The coefficient $\beta_j$ is normalized to -1 in each set of estimates. Sample March 2005 – December 2009. Standard errors in “(.)” parentheses and p-values in “[.]” parentheses.							

The final row of Table 5 performs the same tests with respect to the average price (which includes the Beninois mad Nepalese prices). As in the cases of Kenya and Vietnam, we fail to reject the hypothesis that  $\beta_{Ch} = 0$  so that the average world price is seen as being linked only to the Bangkok price. The estimated  $\beta_{Bk}$  coefficient is 0.75 so prices are seen as varying by 75% of those on the free market. The estimated  $\alpha$  coefficients show that while the Bangkok and Chicago prices both react to the average world price, there is no evidence of any reaction of the world price to the exchange prices.<sup>11</sup>

These results demonstrate that, in the rice market, the prices taken as representing world prices follow prices in local markets round the world rather than *vice versa*. Instead of asking how fast and to what extent changes in world prices are passed through into national

<sup>11</sup> See footnote 8. The same qualification applies here.

prices, we should ask the reverse question of pass-through from national to world prices. The picture we obtain in one of an imperfectly globalized market in which the prices taken as world prices relate to residual transactions and furthermore, these supposedly representative prices do not even move closely together. The Chicago price for rough rice does not appear to have major relevance to world markets. In line with the generally accepted view, the Bangkok spot price appears more important. Nevertheless, that price appears to be reacting to prices in producing and consuming markets rather than determining those prices. In terms of the literature in financial economics, price discovery appears to take place in the producing and consuming markets more than in the centralized spot and futures markets.

## **2.3 Market integration summary**

Maize and rice present a contrasting picture. In both cases, there are two recognized world prices one of which is generated in the Chicago futures markets. In both cases, the two prices relate to different specifications – yellow *versus* white maize and (milled) white *versus* (unmilled) rough rice. The prices therefore differ both in terms of level and monthly changes. However, while the two maize prices are cointegrated and therefore tend back towards a time invariant proportion, there is no clear long term relationship between the two world rice prices. Furthermore, while maize prices do appear to be set on the major international world maize markets, rice prices appear to be determined in a decentralized manner in rice producing and consuming countries.

Standard pass-through models work well for maize. Kenya is seen as having largely insulated itself from changes in world maize prices, but, despite its efforts, Malawi has failed to do so.<sup>12</sup> The same pass-through models work poorly for rice where transmission is largely in the opposite direction from that supposed in those analyses.

## **3 Intra-country price variability**

In this section, I look at the variation of maize and rice prices within the countries under consideration. It would be desirable to do this using household data, as in Kanbur and Grootaert (1994), but these are not available over the time span considered. I therefore consider variation over different geographical localities within each country in order to provide a measure of national price integration.

### **3.1 Methodology**

Countries collect price data on different regional bases – cities, provinces or regions. The degree of regional disaggregation obtainable therefore differs across countries. A near universal problem is that the data often contain large numbers of missing observations. The

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<sup>12</sup> Daviron (2008) stresses “une dynamique propre et endogène des marchés des céréales locales” which results in a “forte instabilité”. The clear local dynamic is evident in our results for Benin, Kenya and Malawi, but the high volatility only in Malawi.

makes measurement of regional variability problematic. The standard procedure of averaging and measuring variability over all available observations is likely to result in a low estimate of price variability over space but may also introduce spurious variability over time. There are two possible ways of dealing with this problem

- Confine the analysis to those regions for which there is complete information.
- Use a statistical model to fill in the missing values.

The former procedure may be sensible if missing observations are confined to a small number of locations. This is generally not the case with the price data I analyze here. There must also be a worry that missing data will relate to smaller and more peripheral locations so that restriction of the analysis to locations with complete data may again result in an under-estimation of variability. This argues for the second, data-filling, strategy.

Let the price of a commodity in region  $j$  and month  $t$  be  $p_{jt}$ , ( $j=1,\dots,n; t=1,\dots,T$ ). Denote as  $S_j$  the set of months for which  $p_{jt}$  is observed for region  $j$ . In order to estimate missing prices, suppose

$$\ln p_{jt} = \ln \pi_t + \delta_{jt} + \varepsilon_{jt}$$

where  $\pi_t$  is the (unobserved) representative national price in month  $t$ ,  $\delta_j$  is the average region  $j$  differential relative to the national average and  $\varepsilon_{jt}$  is a random error. Given estimates  $\hat{\pi}_t$  and  $\hat{\delta}_{jt}$ , an estimate  $\tilde{p}_{jt}$  of a missing price can be obtained as

$$\ln \hat{p}_{jt} = \ln \hat{\pi}_t + \hat{\delta}_{jt} \quad (t \in S_j)$$

The procedure adopted is as follows.

- Provided that there is at least one price observation for month  $t$ , I estimate  $\pi_t$  as the median of the observed prices. The median is preferable to the average in this context since it will be less affected by the pattern of missing observations such as the presence or absence of a low or high price zone.
- In the case that no prices are reported for a particular month  $t$ , which happens only for a small number of isolated months, I interpolate  $\ln \hat{\pi}_t = \frac{1}{2} [\ln \hat{\pi}_{t-1} + \ln \hat{\pi}_{t+1}]$  in two of the countries I discuss.
- This allows estimation of the differentials  $\tilde{\delta}_{jt} = \ln p_{jt} - \ln \hat{\pi}_t$  ( $t \in S_j$ ). Now suppose the differentials are AR(1),  $\delta_{jt} = \kappa_j + \rho_j \delta_{j,t-1} + v_{jt}$ . I estimate the parameters of this AR(1) by OLS over  $S_j$  allowing interpolation of  $\delta_{jt}$  as  $\hat{\delta}_{jt} = \hat{\kappa}_j + \hat{\rho}_j \tilde{\delta}_{j,t-1}$  in the case that  $t-1 \in S_j$  and  $\hat{\delta}_{jt} = \hat{\kappa}_j + \hat{\rho}_j \hat{\delta}_{j,t-1}$  otherwise.
- The national prices analyzed in section 2 are the medians of the monthly prices including the interpolated prices. They are therefore based on an invariant sample of

locations. They will differ slightly from the medians  $\pi_t$  of the reported prices used for in-filling the missing price data.

This results in a balanced panel of price observations.

### 3.2 Maize

Regional information on maize prices is available for the non-Asian countries in our sample.

Benin: The data relate to six locations over the period January 2005 – December 2009.<sup>13</sup>

Kenya: The data relate to six locations over the period January 2005 – December 2009.<sup>14</sup>

Malawi: 26 towns are distinguished over three regions (Central, North and South) but prices relate to the shorter period January 2005 –August 2009.<sup>15</sup> In many towns, data is available for several locations. For example, the data for Dowa in the Central region includes observations on Dowa itself, but also Bowe and Madisi. In such cases, I took the price for the location as the average of the available prices.

Table 6					
Inter-regional maize price variability					
	Coefficient of variation over time			Average coefficient of variation across regions	Correlation (c.v., average)
	Maximum	Average	Minimum		
Benin	33.3%	30.2%	29.1%	19.9%	0.420
Kenya	31.9%	31.7%	33.9%	10.5%	- 0.096
Malawi	50.8%	50.0%	49.8%	14.2%	0.105
Peru	32.6%	15.7%	15.0%	20.4%	0.444

Columns 1-3 give coefficients of variation over time. Column 4 gives the average coefficient of variation over regions. Column 5 gives the correlation of the coefficient of variation over regions with the average price over regions.  
January 2005 – December 2009 (Benin, Kenya and Peru) and August 2009 (Malawi). Missing values have been replaced by estimates – see text.

<sup>13</sup> Locations with number of missing prices (zero if not stated): Azové, Bohicon, Cotonou, Cové (2), Kétou, Pobè. There are 60 observations in total.

<sup>14</sup> Locations with number of missing prices: Eldoret (11), Kisumu (5), Mombasa (3), Nairobi (3), Nakuru (5) and Taveta (23). There are 60 observations in total.

<sup>15</sup> Locations with number of missing prices: Central – Dedza (4), Dowa (7), Kasungu (7), Lilongwe (1), Mchinji (6), Nkhoto (4), Ntcheu (1), Ntchisi (4) and Salima (11); North – Chipita (4), Karonga (4), Mzimba (1), Nkhata (4), Rumphi (4); South – Balaka (5), Blantyre (4), Chikwawa (4), Chiradzulu (5), Machinga (1), Mangochi (4), Mulanje (11), Mwanza (4), Nsanje (1), Phalombe (13), Tyolo (4), Zomba (8). There are 56 observations in total.

Peru: Data are available for 24 of Peru's 25 regions over the period January 2005 – December 2009.<sup>16</sup>

The regional data are summarized in Table 6. The first three columns of the table report the coefficient of variation of respectively the maximum, average and minimum price across regions. Note that these figures will relate to different regions in each month. Prices have shown greatest variability in Malawi and least in Peru. However, in Peru maximum prices have been much more variable than average prices suggesting high volatility in remote areas. (The same is true to a lesser extent in Malawi). The fourth column reports the coefficient of variation across regions, averaged over time. This is a measure of geographical price variability. It is highest for Benin and Peru and lowest for Kenya, consistent with the activities of the stabilizing activities Kenyan NCPB – see section 2.1. The final column reports the correlation between the monthly series on the inter-regional price coefficient of variation and the level of prices. Here we see that in both Benin and Peru, periods of high prices have been associated with greater variability.

Appendix Charts A-1 to A-4 graph the price averages and ranges for the four countries. The presence of a small number of very large outliers has the effect of considerably raising maximum prices in a number of months. It is unclear whether these high recorded prices reflect acute local shortages or are recording errors.

<b>Table 7</b>		
<b>Share of leading two principal components: maize</b>		
	1 <sup>st</sup> Component	2 <sup>nd</sup> Component
Benin	83.5%	8.3%
Kenya	90.5%	3.7%
Malawi	91.8%	3.2%
Peru	63.7%	7.4%
Four national prices	68.9%	15.4%
The table reports the share of the two leading principal components in the logarithmic price levels of regional prices in the four countries. The final row reports the same statistic for the four national prices.		

The question posed at the start of this paper is the extent to which the national average prices may be taken as representative of prices across the entire country. One way of examining this is to look at the share of the variation in local prices accounted for by the leading principal component of these series. This is tabulated in Table 7. For Kenya and Malawi, this share exceeds 90% and it therefore seems completely reasonable to talk of a

<sup>16</sup> The exception is the port city-province of Callao which has an autonomous status. Regions with number of missing prices (zero if not stated): Amazonas, Ancash, Apurimac (23), Arequipa (3), Ayacucho (7), Cajamarca, Cusco (20), Huancavelica (20), Huánuco, Ica, Junín, Lambayeque, La Libertad, Lima, Loreto, Madre de Dios (8), Moquegua (5), Pasco (10), Piura, Puno (41), San Martín, Tacna (17), Tumbes, Ucayali (7). There are 60 observations in total.



national price. The result for Benin is somewhat less than this threshold making it an intermediate case. In Peru, the leading principal component explains less than two thirds of regional prices questioning the representativeness of the national price. This figure is comparable to the share of the leading component in the “world” composed of the four national prices.

The second component reported in Table 7 is, in principle, related to regional variability. I discuss the Peruvian case, where regional variability appears greatest, in section 3.3 in conjunction with the similar results obtained for rice.<sup>17</sup>

### 3.3 Rice

For four of the countries under consideration, Malawi, Nepal, Peru and Vietnam, we have detailed information available on monthly retail prices for rice by region.

Malawi: Prices are available for 26 locations as in maize over the period January 2005 – October 2009.<sup>18</sup>

Nepal Prices are for medium rice, at both the wholesale and retail level, in nine regions over the period April 2005 – March 2009.<sup>19</sup>

Peru: Prices are available for 18 regions over the period January 2005 – December 2009.<sup>20</sup>

Vietnam Prices, which are at the retail level for 25 provinces over the period January 2005 – December 2009. Data for different months were obtained from different sources raising issues of comparability. For seven of these provinces, data are available for less than half the months considered. These were therefore

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<sup>17</sup> In the case of Benin, where prices are reported for only six locations, the second component appears to relate to a single location, Kétou, where the reported prices are unchanged for several months in succession, including over the highly volatile 2007-08 period. The eigenvector weight is 0.92 for Kétou but is negative and around 0.20 for the remaining five locations.

<sup>18</sup> Locations with number of missing prices: Central – Dedza (4), Dowa (4), Kasungu (5), Lilongwe (1), Mchinji (6), Nkhoto (4), Ntcheu (1), Ntchisi (4) and Salima (4); North – Chipita (4), Karonga (4), Mzimba (1), Nkhata (4), Rumphi (4); South – Balaka (5), Blantyre (4), Chikwawa (4), Chiradzulu (2), Machinga (1), Mangochi (4), Mulanje (9), Mwanza (4), Nsanje (2), Phalombe (13), Tyolo (4), Zomba (8). There are 58 observations in total.

<sup>19</sup> Locations with number of missing prices: Banke (3), Dhankuta (9), Jhapa (11), Kailali (12), Kaski (9), Morang (1), Parsa (1), Rupandehi (2). There is a total of 57 months in the sample.

<sup>20</sup> Locations with number of missing prices (zero if not stated): Amazonas, Ancash (37), Arequipa (35), Ayacucho (32), Cajamarca, Cusco (16), Huánuco, Junín, Lambayeque (17), La Libertad (14), Loreto, Madre de Dios (12), Pasco (15), Piura (13), Puno (44), San Martín, Tumbes (16), Ucayali. There are 60 observations in total.



excluded from the analysis as not permitting sufficiently accurate interpolation.<sup>21</sup>

<b>Table 8</b>					
<b>Inter-regional rice price variability</b>					
	Coefficient of variation over time			Average coefficient of variation across regions	Correlation (c.v., average)
	Maximum	Average	Minimum		
Malawi	27.7%	30.9%	35.1%	13.3%	- 0.318
Nepal	28.4%	23.5%	19.5%	16.0%	0.623
Peru	28.2%	22.6%	20.6%	33.0%	0.023
Vietnam	33.8%	33.0%	33.8%	11.2%	0.125
Columns 1-3 give coefficients of variation over time. Column 4 gives the average coefficient of variation over regions. Column 5 gives the correlation of the coefficient of variation over regions with the average price over regions. January 2005 (Nepal: April 2005) – December 2009 (Malawi: October 2009). Missing values have been replaced by estimates – see text.					

Table 8 summarizes the results in the same format as Table 6 for maize.

- The variation in rice prices over time is most acute in Malawi and Vietnam, and least pronounced in Nepal and Peru. Despite this, the Malawian rice price has been much less variable than has been the price of maize, the main staple.
- The Peruvian price data show more considerable variation in maximum than in the minimum prices. The same effect is visible in Nepal but less acutely. The effect is not apparent in Vietnam. Malawi shows the reverse effect with the greatest variability in the minimum price.
- The variability of prices over regions is most pronounced in Peru, and least so in Vietnam.
- In two countries (Nepal and Vietnam), regional price variability appears to have been higher over 2007-08, when average prices were high, than either over 2005-06 or in 2009. In Malawi, by contrast, there was less variability when prices were high.

Appendix Figures A5 - A8 graph the price averages and ranges for the four countries.

As in the discussion of maize prices in section 3.2, we can investigate the extent to which the national average prices may be taken as representative of prices across the entire country. For Malawi, Nepal and Vietnam, the leading principal component of the regional (logarithmic) prices accounts for in excess of 85% of total price variation with the second component accounting for 3% of variation – see Table 9. For these three countries, it does

<sup>21</sup> Locations with number of missing prices: An Giang (4) , Bac Lieu (25), Bac Ninh (21), Ben Tre (23), Ca Mau (29), Can Tho (7), Da Nang (12), Ha Noi (10), Hai Phong (26), Ho Chi Minh (11), Kien Giang (17), Lam Don (13), Lang Son (25), Nam Dinh (20), Nghe An (26), Quang Nin (25), Thai Binh (16) and Tien Giang (0). There is a total of 60 months in the sample. The following seven provinces were excluded on the basis that prices are available for less than 30 of the months considered: Bien Hoa, Binh Duong, Dong Nai, Khanh Hoa, Long Xuyen, My Tho, Nha Trang.

seem reasonable to talk of a national price. In Peru, the leading principal component explains only around two thirds of regional prices questioning the representativeness of the national price. This figure is considerably less than the share of the leading component in the “world” composed of the six national prices.

<b>Table 9</b>		
<b>Share of leading two principal components: rice</b>		
	1 <sup>st</sup> Component	2 <sup>nd</sup> Component
Malawi	88.3%	2.9%
Nepal	90.9%	2.9%
Peru	67.5%	7.5%
Vietnam	89.9%	3.4%
Six national prices	80.0%	11.5%
The table reports the share of the two leading principal component in the logarithmic price levels of regional prices in the four countries for which regional price data are available. The final row reports the same statistic for the six national prices.		

Of the six countries considered in this paper, Peru shows the greatest regional variability for both maize and rice. For maize, I have price data for 24 of Peru’s 25 provinces, while for rice, data extend only for 18 of these regions. For these 18 common regions, the correlation between the eigenvectors defining the weights for the second principal component is 0.733. Since the maize and rice components have been separately estimated, this suggests that they are indeed measuring a common factor.

Analysis of the component weights indicates that the second components are concentrated on five regions for both maize and rice plus three additional regions for maize where no price data is available for rice. These eight regions are in the south and east of the country. Appendix Figure A-9 graphs the premium in the local maize and rice prices averaged across these eight regions (five in the case of rice) relative to the national price calculated as the median across all 24 regions (18 for rice).<sup>22</sup> Rice prices in these five regions were around 30% than the national median from mid-2005 to mid-2006 and again through 2009. Maize prices were over 20% high than the national median in these eight regions in 2009 but did not deviate from the national pattern in 2005-06. These differences are large and underline the problematic nature of the supposedly national price for maize and rice in Peru. However, these problems do not arise in the other five countries considered.

<sup>22</sup> The eight regions are Apurimac (not rice), Ayacucho, Huánuco, Loreto, Madre de Dios, Moquegua (not rice), Pasco, Puno, Tacna (not rice). See Appendix Figure A-10 for a regional map of Peru. Source: <http://mapsof.net/peru/static-maps/png/peru-regions-and-departments>

### **3.4 Regional price variability summary**

The analysis has demonstrated considerable diversity in regional price variability across countries. If a greater number of countries had been considered the extent of diversity would probably have been greater. Nevertheless, some general conclusions may be drawn

- Kenya and Vietnam stand out as having successfully moderated the impact of the variability of world prices on national prices. In Malawi, by contrast, either because of poor policy or because of local weather variability, maize price variability has been considerably higher than world levels.
- Benin and Peru show the most pronounced inter-regional price variability. In both countries, but particularly in Peru, movements in national average price levels have been unrepresentative of prices in the south and east of the country over considerable parts of the period analyzed. In Kenya and Malawi, by contrast, national average prices do appear representative, this despite the high variability of maize prices in Malawi.

There is a general, but not universal, tendency for price variability across regions to be positively associated with the level of prices with the result that the high prices in 2008 were also the most variable on a regional basis. This tendency was apparent in Benin, Nepal, Peru and Vietnam but not in Kenya or Malawi.

## **4 Conclusion**

The 2007-08 food price spike did impact grains prices in the six developing countries considered in this study but the impact was different across the two grains considered and across countries. The commonly quoted Chicago and Johannesburg (SAFEX) maize prices do appear to provide good guides to prices in national and regional centres with the SAFEX price for white maize more important in Africa than outside. The Bangkok and Chicago rice prices, by contrast, provide a much less secure guide to national and regional rice prices and in any case do not move together in a well-defined manner. Policy-makers should therefore beware of attaching weight to these “world” rice prices which relate to residual transactions. Price discovery in rice appears to take place more in producing and consuming countries than in centralized spot and futures markets. Nevertheless, rice prices have tended to rise over the five year period considered.

Pass-through varies across countries. In past, this depends on whether governments have attempted to stabilize prices to domestic consumers and, if so, how successful they have been in doing this. Kenya and Vietnam appear to have been relatively successful and Malawi relatively unsuccessful in this regard. But in any case, national rice prices have been much less variable than the commonly quoted Bangkok prices.

The paper also reports a systematic method for considering price variability at the regional level. The methodological problem here is that of missing price data which can result in

variability being measured across variable numbers of locations. Once this problem has been corrected, one can compare regional variability across countries. This is relatively low in Kenya, Malawi and Vietnam but high in Peru. Indeed, in Peru, national average prices cannot be taken as clearly representative of prices across the entire country.

Finally, the paper has considered the extent to which high prices are also more spatially variable. There appears to be a general but not a universal tendency in this direction. This finding has the troubling implication that in high price periods, such as 2008, national average prices provide a less clear guide to policy than in normal periods. The paper has proposed a novel procedure for analyzing this variability.

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## Appendix: Regional Price Variability Charts

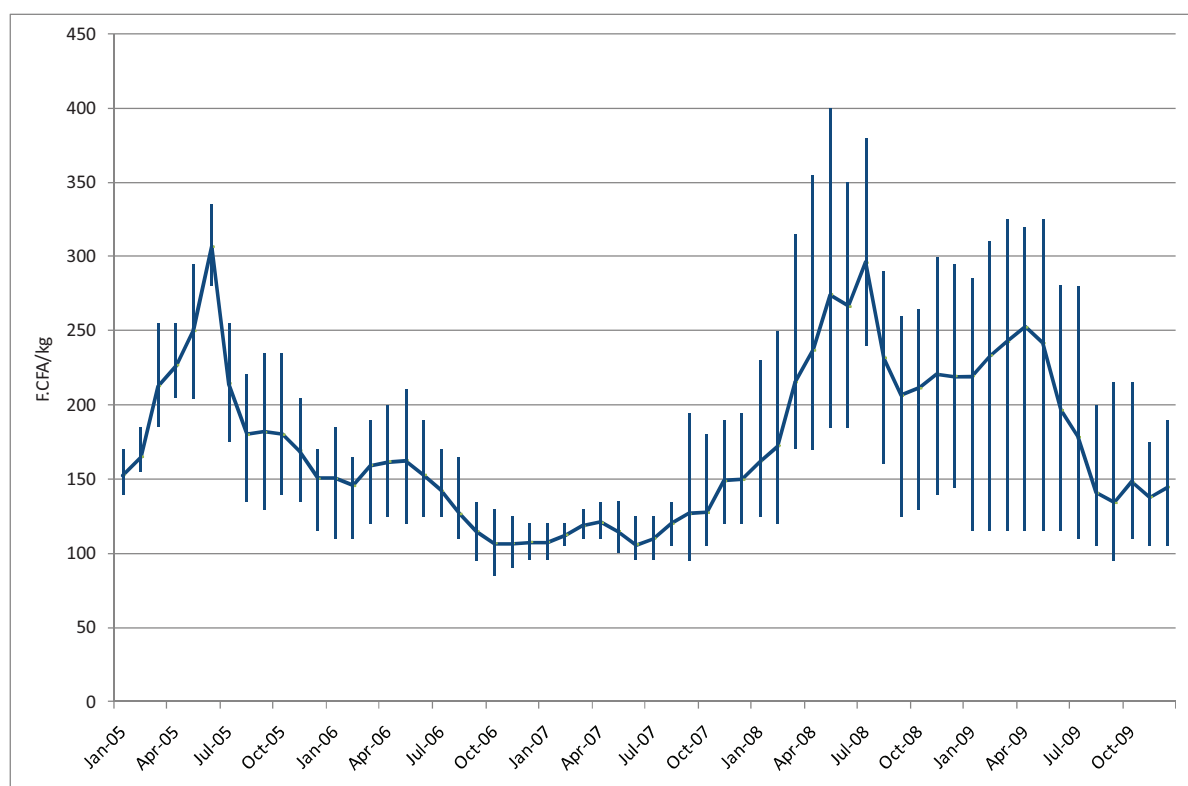
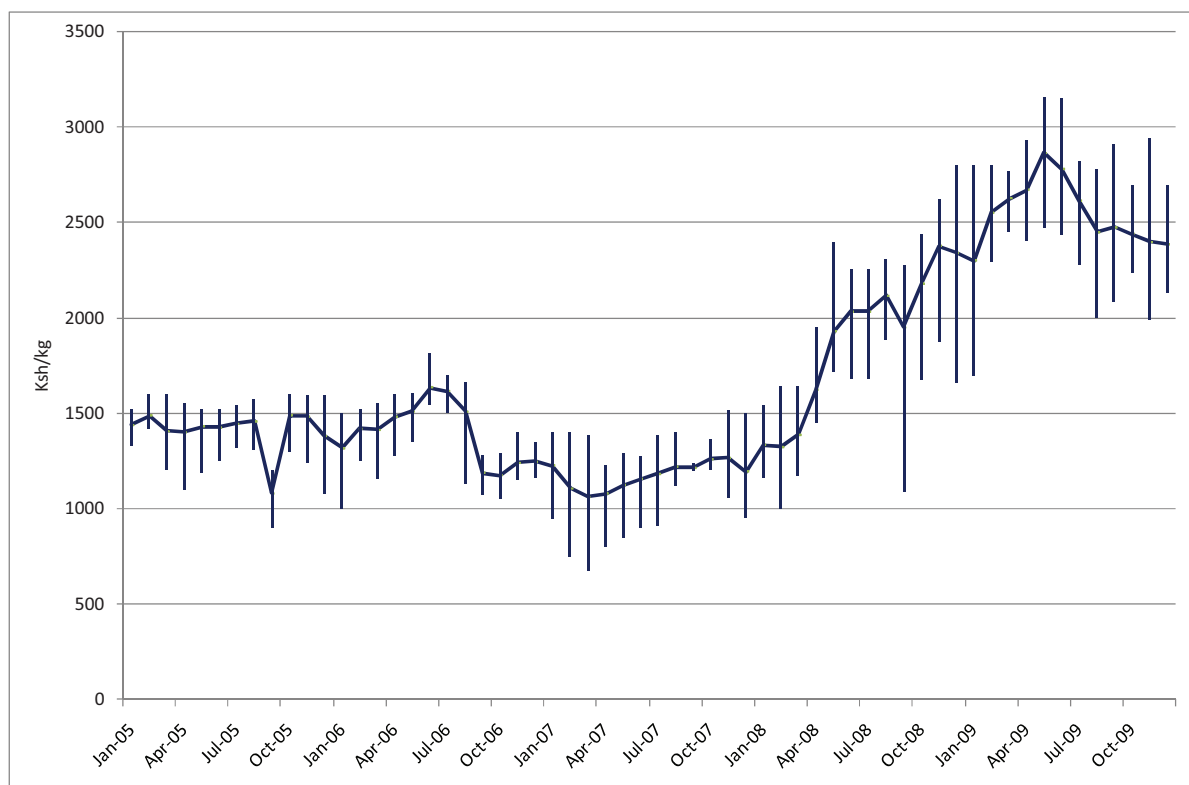
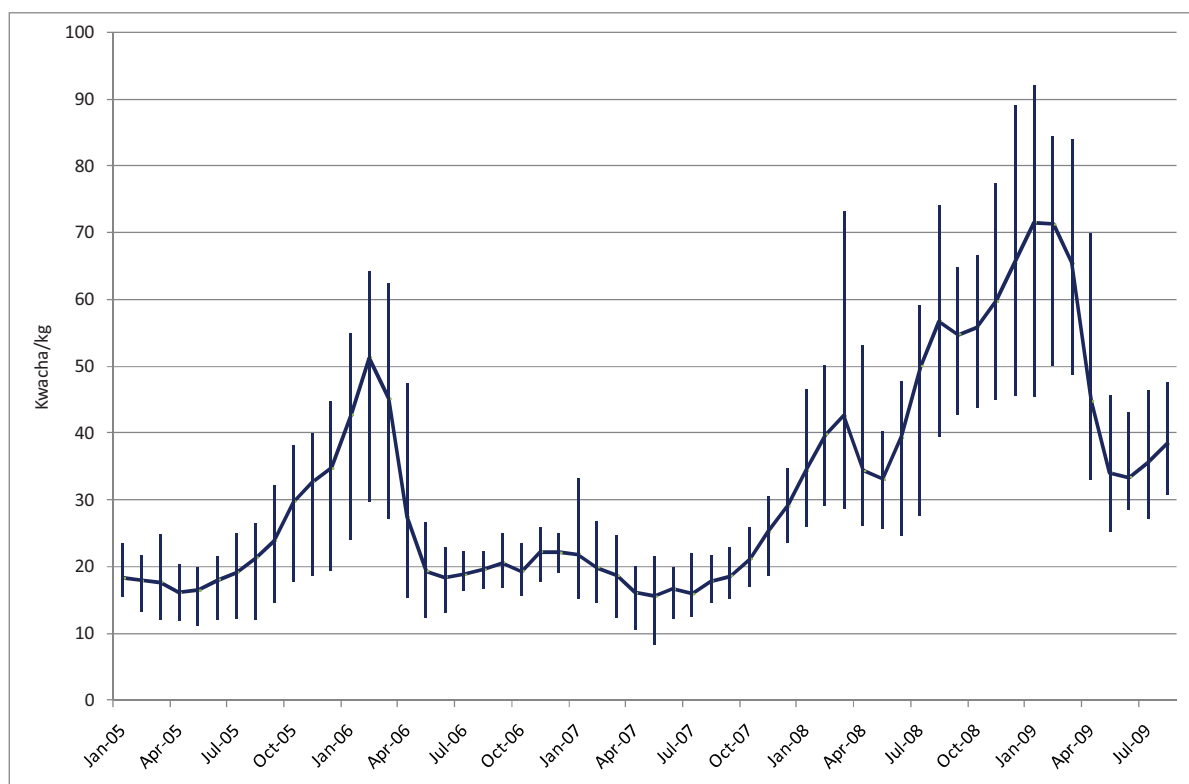


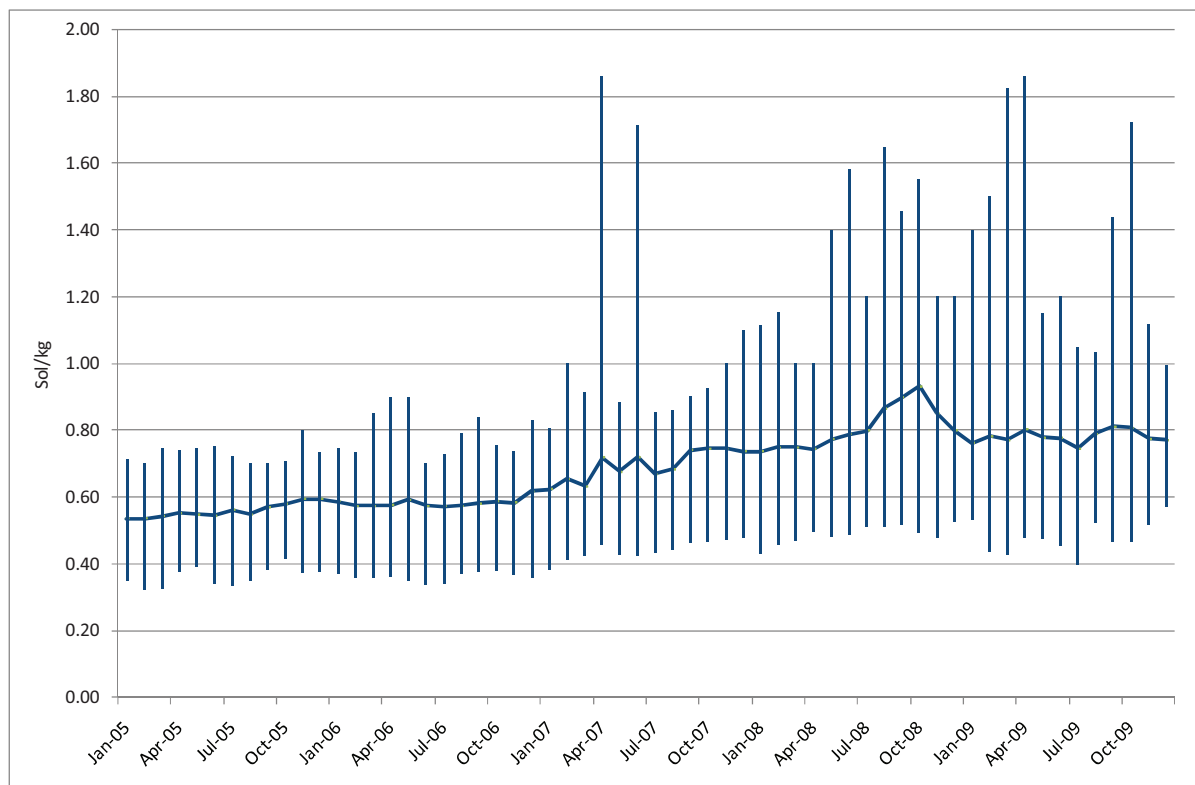
Figure A- 1: Benin, maize



**Figure A- 2: Kenya, maize**

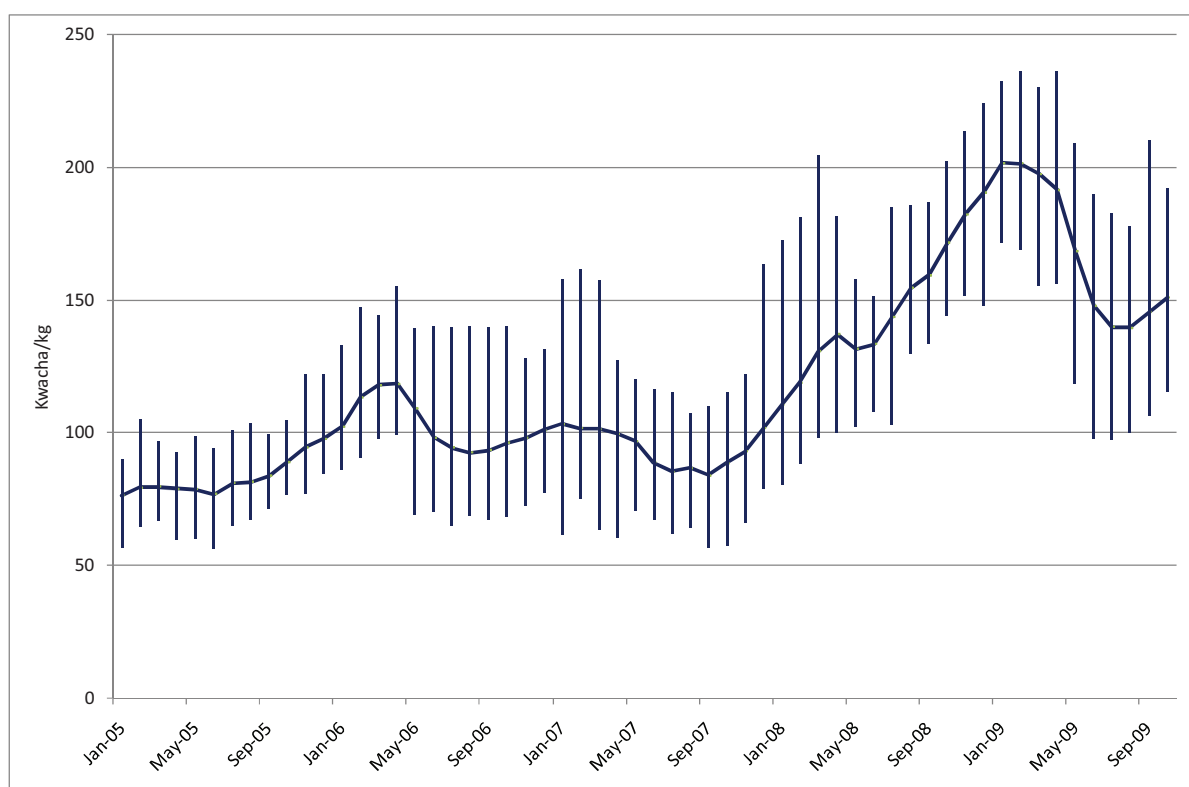


**Figure A- 3: Malawi, maize**

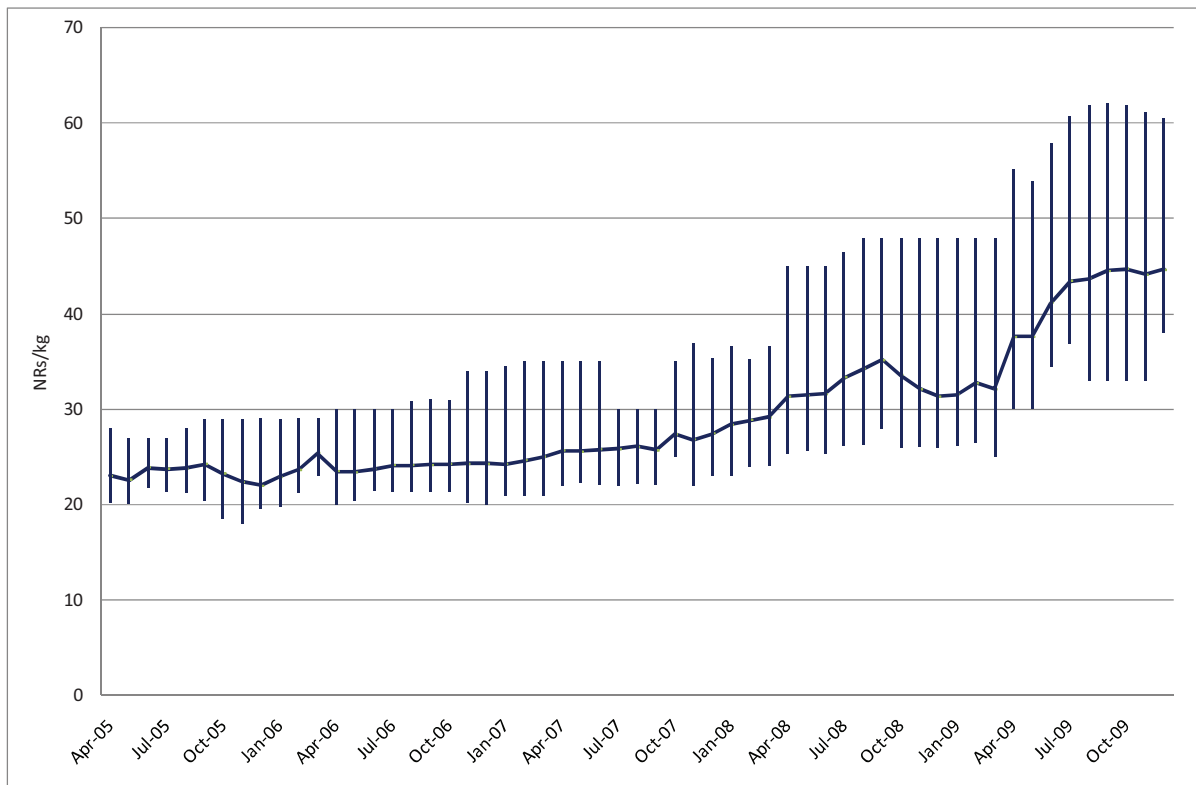


**Figure A- 4: Peru, maize**

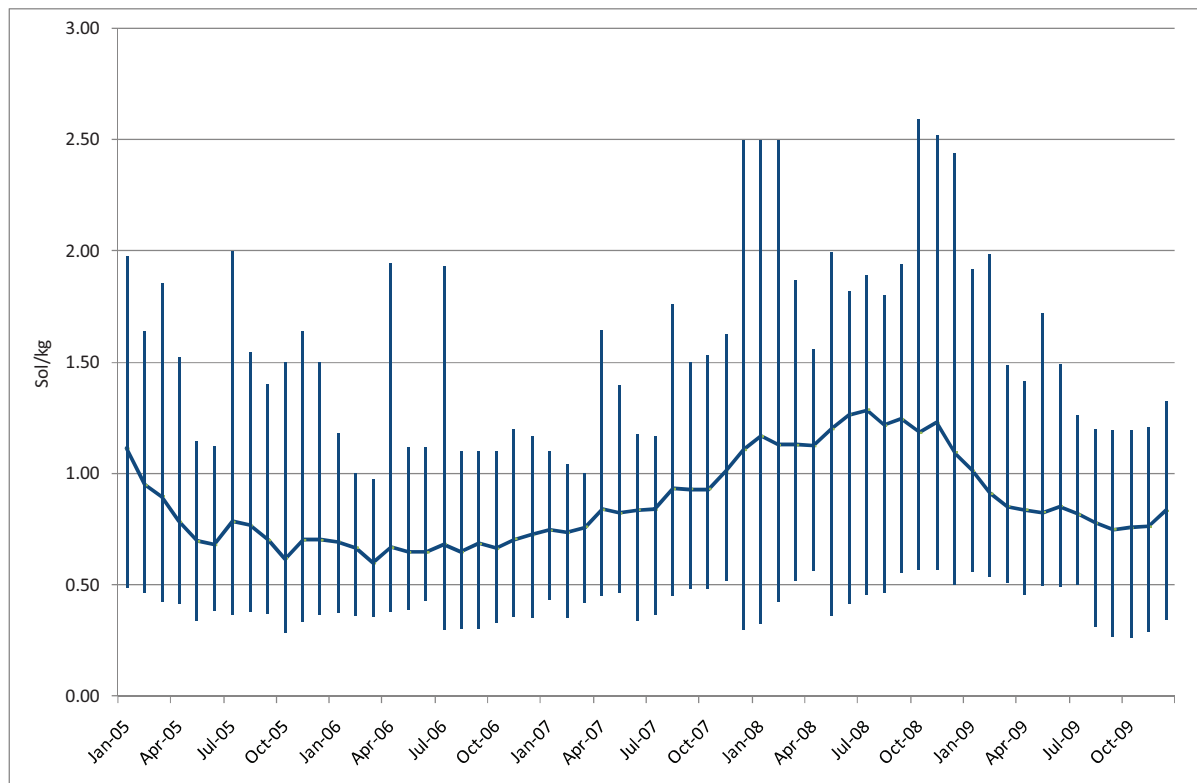




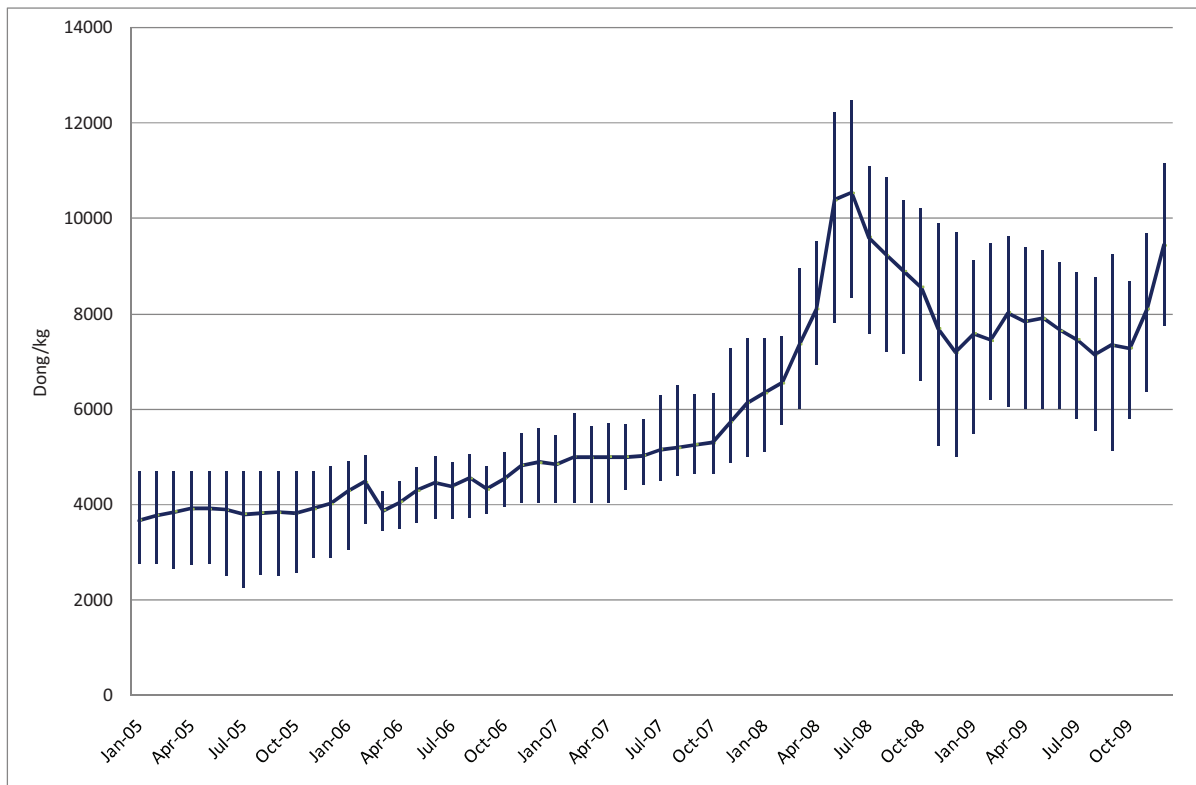
**Figure A-5: Regional variation in rice prices, Malawi, 2005-09**



**Figure A-6: Regional variation in rice prices, Nepal, 2005-09**



**Figure A-7: Regional variation in rice prices, Peru, 2005-09**



**Figure A-8: Regional variation in rice prices, Vietnam, 2005-09**



**Figure A- 9: Premium of high price regions over median, Peru, 2005-09**



**Figure A- 10: Peru, administrative regions**



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University of Trento  
via Inama, 5 - 38122 Trento (IT)

